CEO CAREERS IN REGULATED ENVIRONMENTS: EVIDENCE FROM ELECTRIC AND GAS UTILITIES*

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ABSTRACT

We compare CEOs of electric and gas utility firms to CEOs of unregulated firms. Utility CEOs tend to be older when appointed to office, have less prestigious educational backgrounds, and are more likely to have a legal background. Despite these differences, the evidence also indicates that the likelihood of utility CEO turnover is at least as sensitive to stock performance as the likelihood of turnover among CEOs of unregulated firms. We find no convincing evidence that utility CEOs stay in office longer than their unregulated counterparts, although they are less likely to be overtly forced from office or replaced by an executive from outside the firm. Finally, the evidence suggests that regulatory expertise is valued in the selection of new utility CEOs.

I. INTRODUCTION

The impact of regulation on executive employment contracting has been the subject of considerable research. However, we still have a relatively incomplete picture of how regulation affects managerial labor markets. For example, while recent studies indicate that CEOs of regulated firms are typically paid less than their counterparts at unregulated firms and that the sensitivity of CEO pay to performance is relatively low in regulated environments, the appropriate interpretation of this evidence is subject to debate.[1] One interpretation of these findings is that the political nature of the regulatory process constrains the set of feasible wage packages that can be offered to an executive. An alternative interpretation is that managerial talent and effort are not very important determinants of profits in regulated environments.

Distinguishing between alternative interpretations such as these requires a more complete picture of the labor market for utility managers. Little is known about other dimensions of executive careers in regulated firms, such as the career characteristics of utility managers or the relation between a regulated firm's performance and management turnover.[2] An analysis of these non-wage related issues may help us better understand the actions and objectives of regulated enterprises and shed light on the appropriate interpretation of the compensation evidence. In this study we conduct an analysis of this nature by directly comparing the career

[1] For evidence on the sensitivity of CEO pay to performance see Paul Joskow, Nancy Rose, & Andrea Shepard, Regulatory Constraints on CEO Behavior, Brookings Papers: Micro. No. 1, at 1 (1993); and Paul L. Joskow, Nancy L. Rose & Catherine D. Wolfram, Political Constraints on Executive Compensation: Evidence from the Electric Utility Industry, 27 RAND J. Econ. 165 (1996). For discussions of the interpretation of this evidence see Sam Peltzman, Regulatory Constraints on CEO Compensation. Comments and Discussion, Brookings Papers: Micro. No. 1, at 63-66 (1993); and John R. Meyer, Regulatory Constraints on CEO Compensation. Comments and Discussion, Brookings Papers: Micro. No. 1, at 59-62 (1993).

[2] In contrast, there is an extensive literature examining management turnover for unregulated firms. See, for example, Ann T. Coughlan & Ronald M. Schmidt, Executive Compensation, Management Turnover, and Firm Performance: An Empirical Investigation, 7 J. Acct. and Econ. 43 (1985); Jerold B. Warner, Ross L. Watts, & Karen H. Wruck, Stock Prices and Top Management Changes, 20 J. Fin. Econ. 461 (1988); Michael S. Weisbach, Outside Directors and CEO Turnover, 20 J. Fin. Econ. 431 (1988); Michael C. Jensen & Kevin J. Murphy, Performance, Pay and Top-Management Incentives, 98 J. Pol. Econ. 225 (1990); and Mark R. Huson, Robert Parrino, & Laura T. Starks, Internal Monitoring Mechanisms and CEO Turnover: A Long-term Perspective, 56 J. Fin. 2265 (2001).

characteristics of a sample of CEOs from regulated electric and gas utilities with a benchmark sample of CEOs from unregulated firms.

We find several interesting differences between the backgrounds of the two groups of executives and some equally interesting similarities. Consistent with Palia, we find that CEOs of electric and gas utilities tend to have graduated from less prestigious institutions than their unregulated counterparts.[3] To the extent that school prestige is a proxy for general managerial talent, this evidence is consistent with the notion that overall managerial ability is relatively less important in electric and gas utilities. We also find that executives at electric and gas utilities are much more likely to have a legal background than are executives at unregulated firms, indicating that legal expertise is more highly valued in a regulated setting. In addition, we find that utilities rarely appoint outsiders to the CEO position, but that the outsiders who are appointed CEO by these firms generally have local regulatory expertise. This evidence suggests that regulatory expertise is valued in the selection of utility CEOs and that this expertise is highly local or firm-specific in nature.

When we examine CEO tenure, we find some evidence that, on average, utility CEOs are older than CEOs at unregulated firms when appointed. However, controlling for starting age, we find that there is little difference between the regulated and unregulated firms in the distributions describing the total length of time a CEO serves in office. While this finding may appear inconsistent with evidence reported by Stigler, we contend that it is not inconsistent once the censored nature of career length data is accounted for.[4]

[3] Darius Palia, The Impact of Regulation on CEO Labor Markets, 31 RAND J. Econ. 165 (2000).
[4] George J. Stigler, Legislative Tenure with a Supplement on the Tenure of Business Executives (1980) (unpublished manuscript, on file with the Center for the Study of the Economy and the State, University of Chicago).

An examination of CEO turnover reveals that the overall sensitivity of turnover to stock performance at utilities is at least as great as that at unregulated firms. Poor performance increases the likelihood of utility CEO turnover. Thus, while overall managerial ability may be accorded less weight in the utility CEO selection process, this evidence suggests that utility CEOs are held as accountable for firm performance as their counterparts at unregulated firms.

The rest of the paper is organized as follows. In Section II we discuss the related literature and some relevant hypotheses. Section III describes the sample selection process. In Section IV we examine differences in CEO backgrounds and in Section V we examine differences in CEO tenure. CEO turnover evidence is discussed in Section VI, and Section VII concludes.

II. MANAGERIAL PRODUCTIVITY AND REGULATORY EXPERTISE

A. Managerial Productivity

Several studies have examined the effect of regulation on the level of wages of lower-level employees. The evidence reported in the early studies in this area is mixed.[5] However, more recent studies by Ehrenberg and Rose report strong evidence that employees of New York Telephone Company and regulated trucking firms, respectively, earned above normal wages.[6]

While lower-level jobs at regulated firms may pay high wages, the wages of the highest-level employees in regulated firms appear to be lower than the corresponding wages at

[5] See, for example, Wallace Hendricks, The Effect of Regulation on Collective Bargaining in Electric Utilities, 6 Bell J. Econ. 451 (1975); and Wallace Hendricks, Regulation and Labor Earnings, 8 Bell J. Econ. 483 (1977).

[6] Ronald G. Ehrenberg, The Regulatory Process and Labor Earnings (Academic Press 1979); and Nancy L. Rose, Labor Rent-Sharing and Regulation: Evidence from the Trucking Industry, 95 J. Pol. Econ. 1146 (1987).

unregulated firms.[7] These relatively low wages could make managerial jobs at regulated firms unattractive. Ehrenberg and Palia offer some relevant evidence. Ehrenberg reports extremely low quit rates among the top-level managers of New York Telephone Company, suggesting that these managerial jobs are not unattractive given the managers' alternative opportunities.[8] Palia, on the other hand, reports that executives at regulated electric and gas utilities are less likely than others to have graduated from prestigious colleges, suggesting that these lower paying jobs tend to attract less talented managerial personnel.[9]

One possible explanation for the evidence on utility executive compensation is that managerial talent and/or effort may be less important in regulated firms than in other firms. Consequently, there may be little benefit to paying utility executives high wages or strongly linking their pay to firm performance. We refer to the idea that managerial productivity is less important in regulated utilities as the *Managerial Productivity Hypothesis*.

Given the structure of the electric and gas utility industry, there are good reasons to suspect that the managerial productivity hypothesis may be valid. Prices for the products sold by firms in this industry are traditionally set by state public utility commissions in a way that is supposed to yield a fair rate of return on invested capital.[10] In the rate setting process, utility commissions traditionally allow most costs to be passed on to consumers. In this environment, it is often argued that there is little incentive for utilities to minimize costs or maximize

[7] See Thomas M. Carroll & David H. Ciscel, The Effects of Regulation on Executive Compensation, 64 Rev. Econ. & Stat. 505 (1982); Anup Agrawal, Anil K. Makhija, & Gershon N. Mandelker, Executive Compensation and Corporate Performance in Electric and Gas Utilities, 20 Fin. Mgt. 113 (1991); Joskow, Rose, & Shepard, *supra* note 1; and Joskow, Rose, & Wolfram, *supra* note 1.

[8] Ehrenberg, *supra* note 6.

[9] Palia, *supra* note 3.

[10] See Paul L. Joskow & Roger G. Noll, Regulation in Theory and Practice: An Overview, in Studies in Public Regulation 1 (Gary Fromm ed. 1981); Paul L. Joskow & Richard Schmalensee, Incentive Regulation for Electric Utilities, 4 Yale J. Reg. 1 (1986); and Paul L. Joskow & Nancy L. Rose, The Effects of Economic Regulation, in 2 Handbook of Industrial Organization 1450 (Richard Schmalensee & Robert D. Willig eds. 1989) for surveys of the rate setting process in theory and in practice.

efficiencies, since the benefits of these actions accrue largely to consumers rather than to shareholders.[11] If the profits of electric and gas utilities are fixed by this "cost of service" based regulation, it follows that shareholders may have little incentive to hire highly talented managers or to motivate managers to maximize profits.

Our data allow us to report evidence relevant to this hypothesis. First, as suggested by Palia, if the prestige of the educational institutions a manager attends is a valid proxy for managerial ability, we would expect executives of regulated firms to have graduated from relatively less prestigious institutions.[12] Second, if the benefits realized from replacing managers of poorly performing firms were reduced by the presence of regulation, we would expect to observe a weaker turnover-performance relation at electric and gas utility firms than at unregulated firms.

B. Regulatory Expertise

In addition to assessing the importance of managerial productivity, we are interested in whether utility firms value regulatory expertise in selecting CEOs. Given the direct relation between allowable rates of return and utility profits, the ability of a CEO to effectively deal with

[11] The question of whether utilities have an incentive to operate efficiently has been widely debated in the literature. Some authors, for example, Paul L Joskow, Pricing Decisions of Regulated Firms: A Behavioral Approach, 4 Bell J. Econ. 118 (1973) and Paul L. Joskow, Inflation and Environmental Concern: Structural Change in the Process of Public Utility Regulation, 17 J. Law & Econ. 291 (1974), have argued that "regulatory lag" can result in substantial incentives to lower costs.

[12] Palia, *supra* note 3.

public utility commissions and to successfully navigate the rate setting process is likely to be important.[13] We refer to this possibility as the *Regulatory Expertise Hypothesis*.

It is difficult to directly measure a manager's regulatory expertise. However, we can observe the backgrounds of managers for evidence of the importance of these skills. In particular, to the extent that legal expertise and training are helpful in dealing with legal, political, and public relations issues and allow managers to more effectively deal with politicians and regulators, evidence that utility managers have such skills can be interpreted to reflect their importance.

Additional evidence on the importance of regulatory expertise can be obtained by examining hiring practices in the utility industry. To the extent that regulatory expertise is specific to the geographic area in which a utility operates, we expect to observe a relatively low rate of outside CEO hiring in electric and gas utilities since the pool of qualified outside CEO candidates will be small. Finally, any outsiders who are appointed CEO should possess backgrounds indicating that they bring valuable regulatory expertise to the firms that hire them.

C. Other Studies of Regulation and CEO Careers

In addition to Palia, there are two other relevant studies that examine the impact of regulation on the careers of utility executives.[14] An unpublished study by Stigler analyzes

[13] Dominic M. Geraghty, Strategic Planning Methods and Objectives for Utilities, in Strategic Management and Planning for Electric Utilities 12 (James L. Plummer, Eugene Oatman & Pradeep K. Gupta eds. 1985) discusses how the regulatory process can be influenced through the careful presentation of information. William T. Gormley, Jr., The Politics of Public Utility Regulation (1983) contends that it is not uncommon for utilities to spend in excess of $1 million on presentations at rate case hearings. For discussion of the political incentives in regulation, see the review of the private interest theory of regulation in George L. Priest, The Origins of Utility Regulation and the "Theories of Regulation" Debate, 36 J. Law & Econ. 289 (1993) and also George J. Stigler & Claire Friedland, What Can Regulators Regulate? The Case of Electricity, 5 J. Law & Econ. 1 (1962) and Harold Demsetz, Why Regulate Utilities, 11 J. Law & Econ. 55 (1968).

[14] Palia, *supra* note 3. Also see Stacey R. Kole & Kenneth M. Lehn, Deregulation and the Adaptation of Governance Structure: The Case of the U.S. Airline Industry, 52 J. Fin. Econ. 79 (1999), which examines changes in corporate governance characteristics when the airline industry was deregulated.

survey data on the average tenure of executives at large public utilities versus the largest industrial firms.[15] Some of Stigler's evidence suggests that utility executives have slightly longer tenures, but the samples are small and the differences are not statistically significant. Our large sample enables us to report comprehensive evidence on utility CEO tenures. In addition, as we discuss below, there are some econometric issues in measuring career lengths that the modern literature on duration analysis allows us address.

Geddes examines turnover patterns in the utility industry by estimating logit models of management turnover for a sample of utility firms.[16] He finds that CEO turnover is not sensitive to changes in accounting profits, but is sensitive to changes in consumer electric rates. Our analysis extends that of Geddes in at least three key dimensions. First, we directly compare turnover in utility firms with turnover in non-utility firms. Second, we consider the role of stock returns as a predictor of management turnover. Finally, we report detailed data on CEO turnover events from news articles. These data include, for example, information on whether the CEO leaves to accept a position at another firm and whether the CEO is overtly forced from office.

III. SAMPLE SELECTION

Our sample of CEOs of large public firms is derived from the annual *Forbes* executive compensation surveys from 1971 to 1995. These annual surveys, which have been used extensively in the executive compensation literature, are published in a regular issue of *Forbes* magazine each spring and report the identity of the highest paid executive at approximately 800 large public firms in the U.S. as of the end of the most recent fiscal year. We use turnover announcements reported in the *Wall Street Journal* to confirm the dates that these individuals

[15] Stigler, *supra* note 4. We thank the referee for bringing this study to our attention.
[16] Richard R. Geddes, Ownership, Regulation, and Managerial Monitoring in the Electric Utility Industry, 40 J. Law & Econ. 261 (1997).

hold the CEO title for the utility and benchmark samples described below.

In examining the impact of regulation on labor market characteristics, it is important to isolate the effects of regulation. It is possible that there may be differences in managerial labor markets across industries for reasons unrelated to regulation. For example, using Parrino's industry homogeneity measure, we report below that the electric and gas utility industry is relatively homogeneous (that is, utility firms employ similar production technologies and compete in similar product markets).[17] Parrino finds that the homogeneity of an industry has an important influence on CEO succession patterns and contends that the precision of performance measures is likely to be greater and the costs of hiring an outside CEO is likely to be lower in homogeneous industries than in heterogeneous industries.[18] He also argues that if it is easier to measure CEO performance in homogeneous industries, this may reduce the importance of incentive compensation in aligning manager and shareholder interests.

The preceding considerations suggest that at least some characteristics of managerial employment contracts in regulated firms may be similar to those in unregulated firms that operate in similarly homogeneous environments. To insure that our results are not driven by differences in industry homogeneity, we compare CEOs of regulated electric and gas utilities to a benchmark sample of CEOs from unregulated firms in relatively homogeneous industries. We use the mean partial correlation industry homogeneity proxy constructed by Parrino to identify

[17] Robert Parrino, CEO Turnover and Outside Succession: A Cross-Sectional Analysis, 46 J. Fin. Econ. 165 (1997). This industry characterization is likely to be especially appropriate during the 1971 to 1994 period that we examine. However, recent efforts to deregulate the industry notwithstanding, most utilities still operate in an environment that may be characterized as relatively homogeneous.
[18] *Id.*

the homogeneous industries from which we select this benchmark sample.[19]

We construct our sample by first identifying all CEO turnovers where (1) the turnover can be identified from the name listing in two successive *Forbes* surveys, (2) data for the firm are available in the Standard and Poors' COMPUSTAT database, and (3) the succession announcement is reported in *The Wall Street Journal*.[20] We identify the industry in which each firm competes using the Standard and Poors' COMPUSTAT database. The regulated electric and gas utilities we examine are those with an SIC code of 4911, 4923, 4924, and 4931. The sample that satisfies the three criteria listed above includes 149 CEO successions at utility firms, 277 successions in the financial industry, and 890 successions in the non-utility/non-financial industries. Since characteristics of financial firms differ from other firms in important ways, we exclude financial firm observations from the benchmark sample.[21]

We sort the remaining 890 observations using the homogeneity proxy for the primary industry in which each firm competes and select, as a benchmark sample, the one-third of the 890 observations from the most homogeneous industries as measured by this coefficient. The industries included in the benchmark sample (the benchmark industries) and those that are not

[19] *Id.* Industry definitions in our analysis are based on two-digit SIC codes. Richard N. Clarke, SICs as Delineators of Economic Markets, 62 J. Bus. 17 (1989) finds that the two-digit SIC industry classification captures common industry characteristics as well as more narrow industry definitions. In estimating the mean partial correlation coefficient, we follow the approach used by Parrino. We first regress the monthly returns for each firm included in the CRSP research database against equally weighted return indices for the market and for the industry using data covering the period January 1970 through December 1988. For each industry, we then compute the arithmetic average of the partial correlation coefficients for the industry index, across all industry firms, to obtain the mean partial correlation coefficient for the industry.

[20] The first criterion is not strictly met in all cases as we discovered in our data collection some instances where a CEO is appointed and then relinquishes his position before the next compensation survey is published. These CEOs, which we identify from *Wall Street Journal* articles, are included in some of our summary statistics tables below but are not included in the regression analysis due to insufficient data availability during the CEOs (short) tenure.

[21] See, for example, Joel F. Houston & Christopher M. James, Management and Organizational Changes in Banking: A Comparison of Regulatory Intervention with Private Creditor Auctions in Nonbank Firms, 38 Carnegie-Rochester Conference Series on Public Policy 143 (1993) and Glenn R. Hubbard & Darius Palia, Executive Pay and Performance: Evidence from the U.S. Banking Industry, 39 J. Fin. Econ. 105 (1995) for evidence on differences between financial and non-financial firms.

are reported in Table 1.

IV. CEO BACKGROUNDS

We obtain data for our analysis of CEO backgrounds from the *Forbes* surveys. Because *Forbes* varied its data reporting over the years, some data we use in our analysis of managerial backgrounds are only available for a limited number of years. For example, all *Forbes* surveys up to fiscal 1992 report whether the executives are members of the founding family and classify their career backgrounds. Educational backgrounds, on the other hand, are reported only after fiscal 1985. Since we expect summary statistics on the managerial background variables to change very slowly over time, we report these statistics for the years 1970, 1980, and 1990.

Table 2 shows that considerably fewer CEOs in the utility industry are founding family members than in the benchmark industries in all reported years. This is not surprising since utility firms tend to be older and more established than unregulated non-financial firms. Since firms headed by CEOs who are members of founding families may have very different managerial characteristics than firms not run by founding family members, we exclude founding family CEOs in our calculation of the other summary statistics in Table 2.[22]

The *Forbes* surveys report a variable each year, up to fiscal 1992, indicating a CEOs background as determined by the *Forbes* editors. The background categories include finance, marketing, operations, administration, engineering, and legal. The regulatory expertise hypothesis advanced earlier suggests that utility CEOs are more likely to have legal backgrounds. To investigate this prediction, we compare in Table 2 the fraction of CEOs with legal backgrounds in 1970, 1980, and 1990 for the utility and benchmark industries. As the

[22] This allows us to compare regulated utilities with benchmark industry firms holding founder status fixed. Also note that *Forbes* does not report some of the variables we examine, for example, career background type, for the founder run firms.

figures indicate, the fraction of utility CEOs with legal backgrounds is approximately three times that of the benchmark group and this difference is highly significant in all years. Since the *Forbes* definition of a legal background could be somewhat subjective, we use the education data reported by *Forbes* to assess the fraction, in 1990, of CEOs with legal degrees and find that 28.9 percent of utility CEOs in this year had a law degree. This is significantly higher than the 7.6 percent figure for the benchmark industry sample and is consistent with the regulatory expertise hypothesis.

To investigate whether utility CEOs have less prestigious educations, as suggested by Palia, we collect education data from the 1990 *Forbes* survey.[23] We record whether the CEO holds an undergraduate degree and/or an MBA degree and, for each CEO, we calculate a measure of the prestige of his or her undergraduate (MBA) institution by recording the average SAT (GMAT) score. We calculate these averages using the procedure described by Chevalier and Ellison.[24]

As the figures in Table 2 indicate, the fraction of CEOs with undergraduate or MBA degrees does not differ significantly between the utility and benchmark industries. However, the evidence indicates that utility executives tend to attend less selective institutions. In the case of undergraduate schools, the mean school SAT average of the utility executives is 1,174.81, as compared to 1,228.32 for the benchmark industry group. This 53.51 point difference is

[23] Palia, *supra* note 3.
[24] Judith Chevalier & Glenn Ellison, Are Some Mutual Fund Managers Better than Others? Cross-Sectional Patterns in Behavior and Performance, 54 J. Fin. 875 (1999). Our data on standardized undergraduate test scores are primarily obtained from the Princeton Review Complete Book of Colleges, 1999 edition. When data are not available from this source, we also search the *U.S. News and World Report* Education Web Page and individual school web pages (January 1999). Similar to Chevalier and Ellison, *Id*, in some cases we use data from a school's Princeton Review selectivity index to infer an SAT average. Our primary source for GMAT averages is the *U.S. News and World Report* Education Web pages (January 1999). When data are not available from this source, we use data from Peterson's 1999 Business School Guide or individual school web pages (January, 1999). For many foreign programs standardized test scores are missing. More details on the collection of these data are available from the authors.

significant at the 5 percent level (*t*-statistic of 2.56). Similarly, the mean school GMAT average for the executives with an MBA degree is 614.15 for the utility executives and 654.50 for the benchmark industry sample. The samples are smaller here since only approximately 20 percent of sample CEOs have an MBA degree, but the difference is still significant at the 10 percent level (*t*-statistic of 1.84). This evidence is consistent with the managerial productivity hypothesis to the extent that school SAT and GMAT averages are valid proxies for general managerial talent.

V. CAREER LENGTHS

Before examining CEO turnover, we first examine the distribution describing the total length of time a CEO serves in office by using all executives from the utility and benchmark industries listed in the *Forbes* surveys from 1971 to 1994. We are particularly interested in whether utility CEOs tend to stay in office longer than their counterparts in unregulated industries. To examine this issue, we treat each *Forbes* listed executive as a single observation and ascertain the total time the individual serves as CEO.

In this analysis we identify the year an individual starts his or her CEO service by identifying the year the CEO first appears in the *Forbes* annual survey and then subtracting the number of completed years of CEO service reported in that same annual survey. We can identify a CEO's last year in office in all cases where a firm is listed in two successive annual issues of *Forbes* and the name of the listed CEO changes. However, since some firms exit the *Forbes* surveys over time and our sample period stops at the end of fiscal 1994, we do not observe the end of many executives' tenure as CEO. This raises a censorship issue.[25] To account for this

[25] This is known as "right censorship" in the language of the econometric literature on duration analysis. See, for example, Nicholas M. Kiefer, Economic Duration Data and Hazard Functions, 26 J. Econ. Lit. 646 (1988) and William H. Greene, Econometric Analysis (1993).

censorship, for each executive we note the number of years that we can confirm that he or she is CEO and whether the observation is censored (that is, whether this total is the actual number of years served or a lower bound on the number of years served). Modern econometric techniques for analyzing duration data in the presence of censoring allow us to estimate the (uncensored) underlying distribution of the length of time in office.

In our analysis of career lengths, we examine two sub-samples. The first sub-sample consists of CEOs listed in the first (1971) *Forbes* survey, which corresponds to the end of fiscal 1970. We refer to this sub-sample as the 1970 sub-sample. The second sub-sample is the set of CEOs that enter office after the end of fiscal 1970. We refer to this as the post-1970 sub-sample. This division helps us examine whether there have been any changes over time in the difference between utility CEOs and their counterparts in unregulated industries. We exclude founders in our analysis of CEO tenure.

One variable that is likely to be important in predicting the length of time a CEO is in office is the individual's age when he or she is appointed to the position, since younger executives can serve longer before they reach retirement age. As reported in Table 3 for the 1970 sub-sample, CEOs in unregulated industries are typically 3 years younger than the utility CEOs when appointed (49.68 vs. 52.67 years of age), and this difference is significant at the 1 percent level (*t*-statistic of 2.62). This difference appears to narrow over time, as the difference in the post-1970 sub-sample is small (53.52 vs. 53.91 years of age) and insignificant. Nevertheless, the significant difference in the 1970 sub-sample highlights the importance of controlling for age at appointment when assessing whether utility CEOs are relatively more secure in their positions.

To examine potential differences in career lengths, we report in Table 3 statistics for the distribution describing the length of time utility CEOs and benchmark industry CEOs are in office. We use the Kaplan-Meier product limit estimator that takes into account censoring to estimate the underlying (uncensored) distributions.[26] For the 1970 sub-sample, the estimated distribution of the length of time a CEO serves is fairly similar for the utility group and the non-utility group. The 25th percentile (7 years) and 50th percentile (12 years) of the estimated distributions are the same, while the 75th percentile is greater for the non-utility CEOs (22 years versus 16 years). This suggests that the probability of having a very long CEO career at a utility firm is less than at an unregulated firm, a possibility that we formally test below. For the post 1970 sub-sample, the estimated 25th, 50th, and 75th percentiles of the CEO tenure distribution are all one year greater for the utility group than for the non-utility group. This evidence suggests that utility CEOs may stay in office slightly longer than non-utility executives in the later period, but it not conclusive since it does not control for starting age.

We next estimate a Cox proportional hazard model to formally test whether the typical utility CEO serves for a different time length than his or her counterpart in an unregulated industry.[27] This procedure estimates the effect of explanatory variables on the height of the hazard function describing the total length of time a CEO serves. In this model, a positive coefficient for a variable implies that an increase in that variable proportionally increases the hazard function describing career lengths, thus implying a shorter average career.

The estimated coefficients for this model are presented in Table 4. In Column 1, which reports estimates for the 1970 sub-sample, the coefficient for the utility dummy variable is positive and significant at the 10 percent level, suggesting that utility CEOs tend to spend a

[26] See Greene, *supra* note 26, for a description of the Kaplan-Meier estimator. These estimates are calculated using the "stsum" command of the Stata 5.0 software package.
[27] See Kiefer, *supra* note 26, and Greene, *supra* note 26, for discussions of duration model estimation. These estimates are calculated using the "stcox" command of the Stata 5.0 software package.

shorter time in office than CEOs of the benchmark firms. However, this may largely reflect our previous finding that for the 1970 sub-sample utility CEOs tended to assume the CEO position when they were relatively older. To control for this, in Column 2 we include the age the executive started as CEO as an additional explanatory variable. In this specification the coefficient for the utility dummy variable becomes negative and insignificant, while the age coefficient is positive and highly significant (*t*-statistic of 12.61). These results suggest there is no difference in the expected tenure of two CEOs who are appointed at the same age, regardless of their regulatory status. The results for the post-1970 sub-sample, presented in Columns 3 and 4, also indicate no significant effect of the utility dummy variable on career length. Based on these findings, we conclude that there is no convincing evidence of a difference in expected CEO career lengths between utility and benchmark industry CEOs who begin office at the same age.

VI. CEO TURNOVER

A. Classifying CEO Turnover

The evidence reported above on career lengths suggests that the mechanism governing management turnover may not be significantly affected by regulation. While this evidence is interesting and informative, it is only suggestive in that it does not address the circumstances that lead to a CEO departure. If CEO actions and abilities affect firm performance, then inferior firm performance may lead to CEO turnover.[28] If managerial inputs are less important in regulated than unregulated firms, we expect the turnover-performance relation to be relatively weak for utility CEOs.

We investigate this prediction by directly comparing turnover-performance sensitivities

[28] There is a large literature on management turnover in unregulated firms that documents a negative turnover-performance relation. See, for example, Coughlan & Schmidt, *supra* note 2; Warner, Watts, & Wruck, *supra* note 2; Weisbach, *supra* note 2; Jensen & Murphy, *supra* note 2; and Huson, Parrino, & Starks, *supra* note 2.

between the utility and non-utility firms using the sample of CEO turnovers in the utility (149 turnovers) and benchmark (281 turnovers) industries described in Section III. In this analysis, we use all firm-year observations during the tenure of each departing CEO, beginning the first full year after the CEO is appointed and ending in the year the turnover is announced. This sample of firm-years is augmented with all of the firm-year observations during the tenures of 45 CEOs listed in the 1995 *Forbes* survey that were in office during the entire fiscal 1970 to 1994 period.

For each CEO turnover, we collect data on firm and executive characteristics at the time of the turnover from news articles, corporate proxy statements, Dun and Bradstreet's *Reference Book of Corporate Managements*, and various Marquis *Who's Who* publications.[29] Much of the CEO turnover literature focuses on predicting CEO turnover without regard to the circumstances surrounding it, since determining the exact nature of a CEO succession (forced, voluntary, natural retirement, etc.) can be difficult, as firms are often reluctant to publicly announce the reason for the change. In most of our analysis we also focus on predicting CEO turnover in general, although we do examine the nature of the CEO change in some analyses.

Two special types of CEO turnover that we examine are those where an outsider replaces the incumbent CEO and where the incumbent is overtly forced from office. Outside replacements, which we define as successions where the new CEO has been with the firm for one year or less at the time of the announcement, are relatively straightforward to identify. Identifying forced departures is more difficult. If the *Wall Street Journal* announcement

[29] Some variables are only available for this analysis for a limited number of years due to the high cost of collecting data for these variables for other years. For example, managerial ownership figures are derived from corporate proxy statements. These data are not available in the Q-file microfiche or Disclosure microfiche or electronic proxy collections or in the Spectrum database for years prior to 1978. Proxy data for earlier years are only available from hard-copy sources that are not generally easily accessible.

indicates that a CEO is forced from office or departs because of an unspecified policy difference, we classify the turnover as forced. If the departing CEO is under the age of 60 when he or she departs, we also classify the turnover as forced unless (1) the *Wall Street Journal* announcement reports that the executive departs because of health related reasons or death, (2) the *Wall Street Journal* announces the executive is retiring at least 6 months before the retirement date, or (3) a search of business and trade publications reveals that the executive departs to take a similar position at another firm or departs for previously undisclosed personal or business reasons that are unrelated to the firm's activities.

B. Summary Statistics

Summary statistics for CEO tenure are reported in Table 5. These statistics indicate that the mean and median tenures in office are higher for outgoing utility CEOs than for outgoing benchmark industry CEOs, but this difference is not very informative for two reasons. First, these univariate differences do not condition on founding family status or the age at which the CEO is appointed. In a test that is not reported in Table 5, we restrict the comparison to non-founding family CEOs and find that the difference in means is no longer significant (*p*-value of 0.41). Second, the tenure statistics in Table 5 suffer from the censoring problems discussed earlier since they are calculated only for CEOs who we know left office. If relatively more of the non-utility executives are affected by the censoring, then we may systematically exclude a disproportionate number of long tenure non-utility CEOs in these simple calculations.

The other statistics in Table 5 reveal that stock ownership by the CEO and all officers and directors is significantly lower in the utility sample. The lower stock ownership by CEOs of utility firms is consistent with the weak pay-performance relation reported by Joskow, Rose, and

Shepard.[30] As Table 5 also indicates, the percentage of CEO changes where the departing executive is less than 60 years of age is significantly lower in the utility sample (15.44 percent versus 30.60 percent).[31] However, the percentage of executives that leave at age 65 and the percentage that leave at age 65 or younger are not significantly different across the two groups. Thus, the evidence indicates that younger non-utility CEOs may be more likely to depart than utility CEOs, but that utility CEOs have higher departure rates between the ages of 60 to 64.[32]

From the evidence in Table 5, it is also clear that outside appointments are less common at utility firms (12.8 percent vs. 21.7 percent, p-value of 0.016). One plausible explanation for this finding is that regulatory expertise could be highly local or firm-specific in nature, and thus it may be relatively less feasible to appoint an outside CEO in a regulated utility. To obtain additional insights on the outside successions at sample firms, we examine the backgrounds of the outside CEO hires in our sample.

Of the 19 outsiders appointed CEO by utilities, eight are from the utility industry and 11 are from another industry. Nine of the 11 CEOs from another industry held positions that afforded them the opportunity to develop regulatory human capital in the hiring firm's jurisdiction and, of the remaining two, one was appointed by a utility that was trying to diversify into unregulated businesses. Furthermore, five of the 11 had been on the board of the appointing firm for at least one year at the time of their appointment. As outside directors, these executives are likely to have had an opportunity to develop relationships with regulators in the firm's operating area. Of the remaining six CEOs who were appointed from another industry, three

[30] Joskow, Rose, & Shepard, *supra* note 1.

[31] Note that this comparison does not control for the fact reported earlier that non-utility CEOs tend to be younger when they enter the CEO position.

[32] As Table 5 indicates, utility executives do appear more likely to depart between the ages of 64 and 66. This difference is entirely a reflection of relatively higher departure rates for utility executives at age 64. Since age is measured with some error, some of these retirements at 64 are likely to be typical "normal" retirements at age 65, and some are likely to reflect the fact utility CEO departure rates in the 60-64 age range are relatively high.

came from the telephone communications industry (SIC 4813), an industry also subject to considerable state regulation, and a fourth was previously general counsel to the appointing firm.[33] Taken as a whole, this evidence is broadly consistent with the regulatory expertise hypothesis.

Turning to departures that we classify as forced, Table 5 shows that these departures are less common in the utility industry than in unregulated industries (5.4 percent versus 19.9 percent, *p*-value of 0.000). This could reflect a relative unwillingness by utility boards to fire CEOs, or simply that utilities are more reticent in their explanation for why a manager is departing. Utility firm boards may also initiate CEO departures that are not outright firings. For example, they may "encourage" a manager to retire early via gentle pressure or an enhanced retirement or consulting package. Thus we are hesitant to draw inferences from this forced departure evidence.

B. Logit Models

We estimate logit models, in which the dependent variable takes a value of one if there is a CEO change in a given year and zero otherwise, to examine the effect of firm performance on the likelihood of turnover.[34] Consistent with much of the turnover literature, we use stock returns as our principal measure of firm performance. For each firm-year observation we construct the variable *Return*, which is defined as a firm's 12-month total stock return less the

[33] These turnover events are described in more detail in an appendix that is available from the authors.
[34] We also estimated multinomial logit models, where the dependent variable equals zero if no turnover occurs and other integer values for the various types of turnover events (namely, forced turnover, outside appointment, non-forced/non-outside appointment). The inferences from these more complicated models are very similar to those for binomial logit models, so we report evidence only for the binomial logit models.

industry average return over the same period.[35]

While stock returns include a great deal of information about firm profitability, some have argued that accounting numbers may have more information content than stock returns concerning managerial ability.[36] Thus, as an alternative performance measure, we construct an accounting based variable called ΔROA which we define to be the annual change in the ratio of a firm's earnings before interest and taxes (EBIT) to its book assets less the average change in this ratio in the firm's two-digit SIC industry in the same year. Given the focus on accounting returns in the rate-setting process for utilities, it is possible that this accounting based performance measure is relatively more important for the utility group.

We also use a number of control variables in the logit analysis that are identified in the literature as predictors of CEO turnover. Since older CEOs are more likely to relinquish the position as part of a normal succession process or to retire, we include a dummy variable which takes a value of one if a CEO is 60 years of age or greater and zero otherwise. We also include the number of years the CEO has been in office, since Geddes finds that this variable predicts turnover in utilities even after controlling for age.[37] Finally we use control variables indicating whether the CEO is a founding family member, since founding family members have been found

[35] By subtracting the industry average, we are measuring a firm's performance relative to a similar cohort. Robert Gibbons & Kevin J. Murphy, Relative Performance Evaluation for Chief Executive Officers, 43 Ind. Labor Relns. Rev. 30 (1990) reports evidence supporting the hypothesis that managerial turnover is affected by relative rather than absolute measures of performance. Given the homogeneous nature of the industries in our sample, we expect industry averages to be particularly useful in benchmarking performance.

[36] See, for example, Sherwin Rosen, Contracts and the Market for Executives, in Contract Economics 181 (Lars Werin & Hans Wijkander eds. 1992).

[37] Our control variables are similar to Geddes, *supra* note 16, who also examines CEO changes in the utility industry. Our analysis differs from Geddes in that we do not include a CEO's salary as an explanatory variable since this variable is likely to be endogenous and causally related to the firm performance variables. An additional difference is that we measure performance relative to the industry average, while Geddes makes no such adjustment. Finally, Geddes uses only changes in accounting returns as a measure of firm performance while we also consider stock returns.

to have long tenures, and firm size (natural log of sales), which has also been found to predict the likelihood of turnover.[38]

In Columns 1 and 2 of Table 6 we report results from logit models estimated separately using the benchmark and utility samples, respectively, with *Return* as the performance measure. In both of these models, the coefficient for *Return* is negative and significant at the 5 percent level, indicating that superior stock price performance lowers the probability of CEO turnover in both the utility and benchmark samples. This evidence suggests that utility boards do pay attention to stock returns when making turnover decisions.

In Column 3 we formally test whether the sensitivity of turnover to performance differs between the utility and non-utility firms. We do this by using all observations to estimate the model and interacting *Return* with a utility industry dummy variable (GEDUM). The coefficient for this interaction term is negative and insignificant, indicating that the turnover-performance sensitivity in the utility sample is, if anything, greater than that in the benchmark sample. Furthermore, the insignificant coefficient estimate for the GEDUM variable (*t*-statistic of -0.47) is consistent with our earlier finding of similar predicted CEO career lengths for the utility and unregulated firms.

The last three columns in Table 6 present results from logit models similar to those in the first three columns, but where the accounting variable ΔROA is used in place of *Return* as the performance measure. The results here are similar to those reported in the first three columns, but they are weaker in the sense that the ΔROA variable appears to be less important than the *Return* variable as predictor of turnover for both the utility firms and the benchmark firms. Similar to the models in which we use the *Return* variable, the coefficient for the interaction of

[38] Parrino, *supra* note 18.

the utility industry dummy variable with the ΔROA performance measure in column 6 is negative and insignificant, indicating that turnover-performance sensitivities are no smaller in the utility industry than in the industries from which the benchmark sample is selected. In unreported results we also estimated logit models corresponding to columns 3 and 6 in Table 6, but where all of the non-performance independent variables were interacted with the utility industry dummy variable. In both of these models the interaction of the performance metric with the utility dummy variable remained negative and insignificant. As one would suspect by comparing column 1 with 2 or column 4 with 5, these regressions revealed that the sensitivity of turnover to age and the sensitivity of turnover to sales are significantly greater for the utility group at the 5% confidence level.

Taken as a whole, the evidence in Table 6 is consistent with prior evidence that there is a negative relation between firm performance and the likelihood of turnover. Comparing the first three columns to the latter three, the evidence also suggests that boards rely primarily on stock returns as a measure of managerial performance in governing turnover decisions, at least in homogeneous industries.[39] Most importantly, we find a significant negative relation between turnover and performance in the utility firms, and there is no evidence that the turnover-performance relation is weaker at these firms than at the benchmark industry firms. This evidence casts doubt on the most extreme interpretation of the managerial productivity hypothesis - utility CEO effort and/or talent does not matter - since it suggests that utility boards do consider stock returns when making turnover decisions

[39]Of course these two performance measures should be highly collinear and they both probably contain some marginal information content. Given their high degree of collinearity, it is not very practical or informative to include them both in the same specification. It is worth noting that industry-adjusted stock performance is likely to be more informative in homogeneous industries than in heterogeneous industries as we define them.

D. Additional Investigation

We emphasize the results in Table 6 since the dependent variable used to estimate the models in this table do not depend on a firm's decision to reveal whether a departure is forced. Nevertheless, it is also informative to estimate models in which the dependent variable equals one if the departure is identified as forced and zero otherwise. We do this in Column 1 of Table 7. Consistent with the summary statistics reported in Table 5, the negative and significant coefficient for GEDUM indicates that forced departures are relatively less likely at the utility firms. In addition, the coefficient on the interaction of GEDUM and *Return* is negative and weakly significant (*t*-statistic of -1.86), suggesting that the sensitivity of forced turnover to performance is actually stronger in the utility group. However, given the non-linearity of the logit model and differences in the constant terms for the two groups, a simple interpretation of differences in turnover-performance sensitivities is problematic. When we estimate this forced departure model separately for the utility firms and the non-utility firms (results not reported in tables), we find a strong negative relation between forced turnover and stock return performance for both samples. Consistent with the evidence in Table 6, this evidence indicates that utility boards appear to pay close attention to stock returns when evaluating CEO performance.

Column 2 of Table 7 reports the results for a logit model, using all utility and benchmark firm observations, where the dependent variable equals one if an outsider is appointed CEO and zero otherwise. Consistent with our earlier observations, the significant (*t*-statistic of -2.15) negative coefficient for GEDUM indicates that outside hiring is less common in utility firms. The coefficient for the return variable is positive and insignificant, while the coefficient for the interaction of *Return* with GEDUM is negative and significant (*t*-statistic of -2.44), indicating that outside hiring is relatively more likely in a utility after a period of low stock returns.[40]

One difficulty in comparing turnover-performance sensitivities across samples of firms from different industries is that variation of the performance measure may differ across the two samples.[41] Thus a similarly low stock return may indicate much poorer performance (in a percentile sense) in one group of firms than in another. To ensure that this does not affect our conclusions, we create a performance measure based on deciles of performance. This decile-based variable is created by assigning to each observation a number from one (worst) to ten (best) based on the decile in which the return for an observation falls. The decile cutoff values are chosen separately for the utility observations and for the non-utility observations.

In analysis not reported in a table, we use the return decile variable instead of the raw return variable to estimate logit models corresponding to the specifications in Columns 1 and 2 of Table 6. With this substitution, the coefficient for the return decile variable is negative and insignificant (*t*-statistic of -1.37) for the benchmark sample and negative and significant for the utility sample (*t*-statistic of -2.02). The results for the model where we combine all observations into a single sample and use the return decile performance variable are presented in Column 3 of Table 7. In this model the coefficient for the return decile variable is negative but not quite significant at the 10 percent level (*t*-statistic of -1.38), while the coefficient for the interaction of the return decile variable with GEDUM is negative and insignificant (t-statistic of -0.44). Thus, similar to the results in Table 6, it appears that the likelihood of turnover is at least as sensitive to performance at utility firms as it is at the benchmark firms, even after measuring performance in

[40] This finding is further supported in unreported results where we estimate logit models separately for the utility firms and the non-utility firms.
[41] Joskow, Rose, & Wolfram, *supra* note 1, discuss the relatively low volatility of utility stock returns. In our sample the volatility of the performance measure variables is indeed smaller for the utility sample. Note that the return decile construction we describe below creates a performance variable with the same volatility for both groups of firms.

a uniform way for the two samples.

To complete our analysis, we consider Geddes' hypothesis that increases in electric and gas rates put pressure on utilities to replace management.[42] Using the Moody's public utility manuals, we construct the variable ΔPRICE, which measures the change in residential rates at an individual utility over an individual year.[43] As the evidence in Columns 4 and 5 of Table 7 indicate, this variable is positive but insignificant when included in logit models predicting all CEO changes at utility firms. Furthermore, when the stock return variable is used as the performance measure (Column 4), the estimated coefficient for *Return* is negative and significant at the 10 percent level (*t*-statistic of -1.86). In Column 5, where we use ΔROA as the performance measure, the coefficient for ΔROA is negative and insignificant (*t*-statistic of -0.96). Given the smaller sample used to estimate these models, due to the price data requirement, and the generally weak earlier results on accounting performance, the lack of significance of the coefficient for ΔROA is not surprising. Overall the evidence in Columns 4 and 5 is consistent with our previous finding that turnover is sensitive to stock return performance for utility firms. Controlling for stock return performance, we do not find evidence supporting Geddes' hypothesis that managerial turnover in utilities is sensitive to residential rate changes.[44]

[42] Geddes, *supra* note 16.
[43] Using the approach of Joskow, Rose, & Wolfram, *supra* note 1, we infer electric and gas rates in a given fiscal year by looking at the ratio of residential sales in dollars to residential sales in units (kilowatt hours for electricity and British thermal units for gas). We collect these data from the Moody's public utility manuals from 1970-1994. If electric (gas) revenues exceed gas (electric) revenues, we assume that electricity (gas) is the firm's dominant output. Once we determine the identity of the dominant output, we define ΔPRICE to be the inflation adjusted annual change in the rate charged for this output less the sample-wide average for changes in rates in that same year.
[44] Geddes, *supra* note 16.

VII. CONCLUSION

We examine non-wage characteristics of executive employment in a sample of regulated electric and gas utilities and compare these to a benchmark sample of unregulated firms. The evidence indicates that utility CEOs are less likely to have graduated from a prestigious institution and are more likely to have a legal background. While we find some evidence that utility CEOs are appointed to office at a relatively older age, once we control for age at appointment, we find that the distribution of the expected time a CEO stays in office appears to be unaffected by regulatory status. Consistent with this evidence on career lengths, we find that the overall rate of management turnover and the sensitivity of turnover to performance are similar for the utility sample and the benchmark sample. Finally, we find that utility CEOs are relatively less likely to be overtly forced from office or replaced by an outsider.

Combining our evidence with the prior evidence on low pay and low pay-performance sensitivities in utilities provides a fairly complete picture of managerial employment in a highly regulated setting. It is interesting that turnover-performance sensitivities in utilities are similar to unregulated firms while pay-performance sensitivities are typically quite low in the utility sector. Using the logic of Joskow, Rose, and Shepard, one explanation for these findings is that pay contracts may be constrained in the utility sector because of the highly political nature of the regulatory process.[45] Thus, our findings are consistent with the idea that managerial inputs are as important in the utility industry as they are elsewhere, its just that you cannot pay these managers in an optimal fashion even though you may be able to remove them from office optimally.

An alternative, and in our view more plausible and balanced interpretation for the entire body of evidence on utility CEOs, is as follows. Given the relatively low wages offered to

[45] Joskow, Rose, & Shepard, *supra* note 1.

executives in this industry, arising from political constraints and/or the diminished role of managers in the creation of profits, the utility industry attracts a different pool of managerial talent. This pool tends to have a lower level of general managerial ability, as measured by educational prestige. While the industry may not attract budding "superstars", managers are not totally inconsequential to profits in these firms and there is a certain skill set, which we refer to as regulatory expertise, which is highly desired by the industry. In the pursuit of managers with these skills, utilities tend to hire CEOs with a legal background. Additionally, owing to the fact that regulatory expertise is mostly local in nature, utilities are more likely to appoint insiders to the CEO position. When they do hire CEOs from outside the firm, they tend to hire individuals with some local regulatory expertise. Finally, similar to unregulated firms, utility boards use firm performance as an indication of a CEO's performance and take actions to encourage poorly performing managers to depart.

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TABLE 1

INDUSTRY HOMOGENIETY MEASURES

SIC Code	Industry	Industry Homogeneity Measure	Number of Turnovers in Original Sample
Panel A: Utility industry and homogeneous (benchmark) industries			
10	Metal mining	0.5276	7
13	Oil and gas extraction	0.4483	12
45	Air transportation	0.4433	18
29	Petroleum refining & related	0.4371	52
49	Electric, gas, & sanitary services	0.4135	149
40	Railroad transportation	0.4113	18
15	Building construction – general	0.3827	7
70	Hotels	0.3679	3
54	Food stores	0.3278	30
26	Paper & allied products	0.3247	41
53	General merchandise stores	0.3021	33
23	Apparel	0.2907	4
78	Motion pictures	0.2879	1
59	Miscellaneous retail	0.2864	4
33	Primary metal industries	0.2855	45
24	Lumber & wood products	0.2848	6
Panel B: Non-homogenous (non-benchmark) industries			
28	Chemicals & allied products	0.2798	104
56	Apparel & accessory stores	0.2793	7
22	Textile mill products	0.2636	4
73	Business services	0.2627	6
65	Real estate	0.2596	1

27	Printing, publishing & allied	0.2568	19
58	Eating & drinking places	0.2534	3
38	Measuring, analyzing & control. Instruments	0.2505	34
80	Health services	0.2448	5
48	Communications	0.2446	32
20	Food & kindred products	0.2386	62
30	Rubber & misc. plastic products	0.2252	13
36	Electronic & other elec. equip.	0.2193	43
50	Wholesale trade – durable goods	0.2185	3
32	Stone, clay, glass & concrete prod	0.2110	23
39	Misc. manufacturing	0.2100	2
35	Industrial & commercial mach. & computers	0.2023	67
37	Transportation equipment	0.1932	66
51	Wholesale trade – nondurable goods	0.1894	21
34	Fabricated metal products	0.1522	14

Note.- Data are reported for non-financial two-digit Standard Industrial Classification (SIC) industries for which sufficient data are available to estimate industry homogeneity proxies. Industries are ranked in order of decreasing homogeneity. The industry homogeneity proxy equals the average among all public firms in each two-digit SIC industry of the partial correlations of the firms' stock returns with an industry index in a two-factor model that also includes a market index. The number of turnovers reported is the total number of CEO changes in an industry identified by following the *Forbes* executive compensation surveys from 1971 to 1995 using the approach described in the text.

TABLE 2

MANAGERIAL BACKGROUNDS: UTILITY VS. BENCHMARK FIRMS

	Mean for Electric and Gas Utilities	Mean for Benchmark Firms	t-Statistic for Difference
Founder (1970)	0.000	0.097	-3.92**
Founder (1980)	0.000	0.130	-5.01**
Founder (1990)	0.000	0.098	-3.93**
Legal background (1970)	0.311	0.122	2.85**
Legal background (1980)	0.298	0.122	2.63*
Legal background (1990)	0.227	0.062	2.94**
Law degree	0.289	0.076	3.45**
Undergraduate degree	0.985	0.946	1.56
MBA degree	0.212	0.178	0.58
Avg. SAT score of undergraduate school	1,174.81	1,228.32	-2.56*
Avg. GMAT of MBA granting institution	616.15	654.50	-1.84+

Note.- These statistics are for a set of executives listed in the *Forbes* executive compensation surveys from firms for which financial data are also available on COMPUSTAT. The electric and gas utilities are firms in SIC industries 4911, 4923, 4924, or 4931. The benchmark firms are all firms in the benchmark industries as indicated by a high homogeneity measure. All measures are derived from the data reported by *Forbes* and are from the 1990 survey unless a different year is indicated. All variables except the founder variable are calculated over the set of all firms at which the CEO is not a member of the founding family. All variables except the SAT and GMAT variables are binary variables with values of 0 or 1. The algorithms for calculating the SAT average and GMAT average are described in the text. The t-statistics are for tests that the differences in the mean values for the electric and gas utilities and the benchmark firms are equal, assuming the two populations have unequal variances.

$^{+}$ $P < .10$; * $P < .05$, ** $P < .01$.

TABLE 3

CAREER CHARACTERISTICS: UTILITY VS. BENCHMARK FIRMS

	Electric and Gas Utilities	Benchmark Firms	t-statistic for Difference
Sub-sample 1: CEOs in office in 1970 (n=192)			
Mean age when appointed CEO	52.67	49.68	2.62**
Mean years with firm when appointed CEO	22.98	20.47	1.29
Length of tenure as CEO – 25th percentile estimate	7 years	7 years	
Length of tenure as CEO – 50th percentile estimate	12 years	12 years	
Length of tenure as CEO – 75th percentile estimate	16 years	22 years	
Sub-sample 2: CEOs starting in office after 1970 (n=499)			
Mean age when appointed CEO	53.91	53.52	0.50
Mean years with firm when appointed CEO	21.34	18.94	1.91+
Length of tenure as CEO – 25th percentile estimate	6 years	5 years	
Length of tenure as CEO – 50th percentile estimate	9 years	7 years	
Length of tenure as CEO – 75th percentile estimate	13 years	12 years	

Note.- These statistics are for the sample of all CEOs listed in the *Forbes* executive compensation surveys for fiscal 1970 through fiscal 1994, except for CEOs who are members of the founding family. The electric and gas firms are in SIC codes 4911, 4923, 4924, or 4931. The benchmark firms are all firms in the benchmark industries as indicated by a high homogeneity measure. Sub-sample 1 consists of the CEOs listed in *Forbes* at the end of 1970. Sub-sample 2 consists of executives that become CEO at some point after the end of fiscal 1970. The age of a CEO when appointed to that position and the number of years the CEO has with the firm when appointed CEO are inferred from the *Forbes* data on CEO tenure, tenure with the firm, and age as of the end of the reported fiscal year. The 25th, 50th, and 75th percentile estimates of the length of tenure as CEO are point estimates of the distribution describing the total number of years an executive serves as CEO measured from the year they start as CEO until the year they step down as CEO. These point estimates are derived from the Kaplan-Meier product-limit estimator of the survivor function, where each CEO is treated as a single observation. These estimates take into account the right-censoring of the data when we do not observe an executive's last year as CEO. This censoring arises when *Forbes* stops following a firm or when an executive is still in office as of the end of the sample period. The t-statistics are for tests that the differences in the mean values for the electric and gas utilities and the benchmark firms are equal, assuming the two populations have unequal variances.

$^{+}$ $P < .10$; * $P < .05$, ** $P < .01$.

TABLE 4
THE EFFECT OF REGULATION ON TENURE AS CEO
PROPORTIONAL HAZARD MODEL ESTIMATION

	(1)	(2)	(3)	(4)
Utility dummy	0.304^{+}	-0.285	-0.153	-0.087
	(0.172)	(0.185)	(0.133)	(0.133)
Age at start of tenure as CEO		0.160^{**}		0.152^{**}
		(0.013)		(0.014)
Sample	CEOs in office in 1970	CEOs in office in 1970	CEOs entering office after 1970	CEOs entering office after 1970
Sample size	192	191	499	499
Log Likelihood	-667.01	-590.39	-1293.31	-1227.63

Note.- The reported coefficients are estimates from the Cox proportional hazard model. Standard errors are reported in parentheses. Each CEO's career is treated as a single observation and the dependent variable is the total number of years an executive serves as CEO. The estimation procedure takes into account the right-censoring of the data where we do not observe an executive's last year as CEO. This censoring arises when *Forbes* stops following a firm or when an executive is still in office as of the end of the sample period. The sub-samples used in each column are described in Table 3 and the text. The utility dummy variable has a value of 1 for the gas and electric utility firms and 0 for the other firms.

$^{+}$ $P < .10$; * $P < .05$, ** $P < .01$.

TABLE 5

SUMMARY STATISTICS FOR CEO SUCCESSIONS

	Electric and Gas Utilities	Benchmark Firms	p-value for Difference
Mean tenure	9.33	8.39	0.098 (t=1.658)+
Median tenure	8.58	6.83	0.005 (z=2.837)**
Mean CEO stock ownership	0.05%	1.25%	0.000 (t=-4.276)**
Median CEO stock ownership	0.02%	0.20%	0.000 (z=-9.456)**
Mean ownership by officers and directors	0.30%	7.22%	0.000 (t=-6.681)**
Median ownership by officers and directors	0.13%	2.06%	0.000 (z=-10.934)**
Percentage of departing CEOs less than 60 years of age	15.44%	30.60%	0.000 (t=-3.745)**
Percentage of departing CEOs 65 years of age	26.85%	23.13%	0.402 (t=0.839)
Percentage of departing CEOs with age less than or equal 65 years of age	84.56%	85.41%	0.817 (t=-0.232)
Percentage of departing CEOs with age between 64 and 66	51.68%	37.72%	0.006 (t=2.777)**

Number (percentage) of successions in which an outsider is appointed	19 (12.8%)	61 (21.7%)	0.016 (t=-2.430)*
Number (percentage) of successions in which new CEO is an outsider who was already a director	5 (3.4%)	9 (3.2%)	0.933 (t=0.085)
Number (percentage) of successions in which new CEO is an outsider from another industry	11 (7.38%)	25 (8.9%)	0.581 (t=-0.552)
Number (percentage) of successions classified as forced	8 (5.4%)	56 (19.9%)	0.000 (t=-4.818)**

Note.- Statistics are for a sample of 149 CEO successions at electric and gas utilities (SIC codes 4911, 4923, 4924, or 4931) and 281 successions at benchmark firms occurring during the 1971 to 1994 period. This sample is chosen from the set of firms profiled in the *Forbes* magazine annual executive compensation survey as described in the text. The tenure figures measure years of service as CEO of the departing executive as of the date of the succession announcement. The ownership figures and board characteristics are collected from proxy statements immediately preceding the succession announcement. Proxy data are available for approximately 2/3 of the succession events. An outside appointment is a succession where the new CEO has been employed by the firm for one year or less at the time of the succession announcement. CEOs who vacate the position prior to age 60 and who do not leave for other employment, for a previously announced retirement, or for health reasons are classified as having been forced out, as are all CEOs where the *Wall Street Journal* indicates that the departing executive was forced from office. Differences between means and percentages are tested using simple two-tailed t-tests assuming unequal variances. Differences in medians are tested using the Mann-Whitney two-sided U test.

$^{+}$ $P < .10$; * $P < .05$, ** $P < .01$.

TABLE 6

LOGIT MODELS OF MANAGEMENT TURNOVER: UTILITIES VS. BENCHMARK FIRMS

	(1)	(2)	(3)	(4)	(5)	(6)
Return	-0.618*	-1.357*	-0.657*			
	(0.293)	(0.633)	(0.299)			
Return *			-0.594			
GEDUM			(0.692)			
ΔROA				-0.897	-5.28^{+}	-0.908
				(0.886)	(3.06)	(0.885)
ΔROA*						-4.22
GEDUM						(3.24)
GEDUM			-0.064			-0.047
			(0.135)			(0.132)
Age dummy	1.08**	1.81**	1.33**	1.08**	1.79**	1.32**
	(0.161)	(0.267)	(0.136)	(0.161)	(0.266)	(0.136)
CEO tenure	.031*	.065**	.041**	.030*	.065**	.040**
	(0.013)	(0.019)	(0.010)	(0.013)	(0.019)	(0.010)
Founding family dummy	-1.35**	-1.75	-1.46**	-1.35**	-1.80	-1.46**
	(0.266)	(1.10)	(0.256)	(0.266)	(1.11)	(0.256)
Log of sales	0.138**	0.368**	0.198**	0.139**	0.368**	0.199**
	(0.049)	(0.085)	(0.043)	(0.049)	(0.085)	(0.042)

Constant	-3.72**	-6.72**	-4.45**	-3.70**	-6.22**	-4.42**
	(0.405)	(0.638)	(0.358)	(0.403)	(0.633)	(0.355)
Sample	Benchmark Firms	Electric & Gas Firms	All Firms	Benchmark Firms	Electric & Gas Firms	All Firms
Number of firm-years	1908	1219	3127	1908	1219	3127
Log Likelihood	-644.72	-364.67	-1018.21	-646.50	-365.49	-1020.81

Note.- The dependent variable equals 1 when a firm experiences a change in CEO during the year and 0 otherwise. Asymptotic standard errors are reported in parentheses under the coefficient estimates. The sample is derived from a set of 149 CEO succession events in electric and gas utilities (SIC codes 4911, 4923, 4924, or 4931) and 281 succession events in homogenous benchmark firms occurring during the 1971 to 1994 period where all sample CEOs are profiled in the *Forbes* magazine annual compensation surveys. The Return variable is an industry adjusted 12-month stock return defined as the total return on the firm's common stock less the average total return of firms in the same two-digit SIC industry. The GEDUM variable is a dummy variable with a value of 1 for electric and gas firms and 0 otherwise. The ΔROA variable is an industry adjusted annual change in ROA, where ROA is measured as EBIT divided by book assets and ΔROA is the firm's annual change in ROA less the industry average of this change. The age dummy equals 1 for CEOs aged 60 and older and 0 otherwise. CEO tenure is measured as the number of years the CEO has been in office in the firm-year. The founding family dummy equals 1 if the CEO is a member of the family that founded the firm and 0 otherwise.

$^{+}$ $P < .10$; * $P < .05$, ** $P < .01$.

TABLE 7

LOGIT MODELS OF MANAGEMENT TURNOVER: ALTERNATIVE MODELS

	(1)	**(2)**	**(3)**	**(4)**	**(5)**
Return	2.46^{**}	.061		-1.46^{+}	
	(.736)	(.586)		(0.782)	
Return *	-3.49^{+}	-3.71^{*}			
GEDUM	(1.87)	(1.52)			
Return decile			-.036		
			(0.026)		
Return decile*			-.018		
GEDUM			(0.041)		
ΔROA					-3.38
					(3.54)
ΔPRICE				1.18	1.28
				(1.55)	(1.55)
GEDUM	-1.42^{**}	$-.753^{*}$	.047		
	(0.544)	(0.350)	(0.253)		
Age dummy	$-.557^{+}$	$.689^{*}$	1.32^{**}	1.94^{**}	1.92^{**}
	(0.329)	(0.286)	(0.136)	(0.306)	(0.306)
CEO tenure	-.0009	-.0000	$.041^{**}$	$.084^{**}$	$.083^{**}$
	(0.0346)	(0.0247)	(0.010)	(0.022)	(0.022)
Founding family		-.616	-1.46^{**}	-1.52	-1.53
dummy		(0.494)	(0.256)	(1.14)	(1.14)

Log of sales	.215*	.066	.200**	.137	.142
	(0.108)	(0.094)	(0.043)	(0.105)	(0.104)
Constant	-5.31**	-4.53**	-4.23**	-4.86**	-4.82**
	(0.916)	(0.749)	(0.378)	(0.772)	(1.55)
Sample	All Firms	All Firms	All Firms	Electric & Gas Firms	Electric & Gas Firms
Dependent variable	Forced CEO changes	Outside Hires	All CEO changes	All CEO changes	All CEO changes
Number of firm-years	2663	3127	3127	851	851
Log Likelihood	-210.27	-286.91	-1019.99	-281.87	-283.14

Note.- In column 1 the dependent variable equals 1 if there is a forced CEO change during the year and 0 otherwise. CEOs who vacate the position prior to age 60 and who do not leave for other employment, for a previously announced retirement, or for health reasons are classified as having been forced out, as are all CEOs where the *Wall Street Journal* indicates that the departing executive was forced from office. In column 2 the dependent variable equals 1 if an outsider is appointed CEO and 0 otherwise. An outsider is defined as a new CEO who is employed by the firm for one year or less at the time of the succession announcement. The dependent variable in columns 3, 4, and 5 equals 1 when a firm experiences a change in CEO and 0 otherwise. Asymptotic standard errors are reported in parentheses under the coefficient estimates. The Return decile variable is constructed by assigning firm-year observations a number from 1 (worst) to 10 (best) based on the decile to which their Return variable belongs. The decile cutoff points are calculated separately for the utilities and for the benchmark firms. The ΔPRICE variable is a measure of the change in electric and gas rates for a utility (see text for details) over the course of the year less a measure of the average change in rates across all utilities in that year. All other variable definitions are the same as reported in Table 6. The founding family variable is excluded in column 1 because a coefficient for this variable cannot be estimated since in every instance where this variable is 1 the dependent variable is 0.

$^{+}$ $P < .10$; * $P < .05$, ** $P < .01$.